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Working Paper No. 0605

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The Effect of Income on Positive and Negative Subjective Well-Being*

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March 2006

Abstract

Increasing evidence from the empirical economic and psychological literature suggests that positive and negative well-being are more than opposite ends of the same phenomenon. Two separate measures of the dependent variable may be needed when analyzing the determinants of subjective well-being. We argue that this conclusion reflects in part the use of too restrictive econometric models. A flexible multiple-index ordered probit panel data model with varying thresholds can identify response asymmetries in single-item measures of subjective well-being. An application to data from the German Socio-Economic Panel for 1984-2004 shows that income has only a minor effect on positive subjective well-being but a large effect on negative well-being.

JEL Classification: I31, D12, C23
Keywords: generalized ordered probit model, marginal probability effects, random and fixed effects, life-satisfaction.

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1 Introduction

For economists, the relationship between material well-being and overall quality of life is a matter of great concern. Due to the lack of good life-quality measurements the two are often equated, leading to prescriptions of policies and institutions that focus on income only. Fortunately, much progress has been made in recent years in measuring quality of life, based on the principle that individuals are the best judges of their own well-being, and that their judgment can be elicited using survey-based indicators of subjective well-being (SWB). This approach uses individual responses to a question such as "How satisfied are you with your life, all things considered?" with categorical answers ranging from "completely dissatisfied" to "completely satisfied". Self-reported life-satisfaction has been shown to be a reliable, valid, and consistent measure of individual well-being (see Diener and Lucas 2000, and the references therein).

Using this widely available new instrument, the relationship between material well-being and quality of life can be and has been subject to rigorous empirical analysis. There is by now a large literature on the relationship between income and SWB, with data for different countries, different points in time, and using different model specifications. Frey and Stutzer (2002), Layard (2005), and Diener and Biswas-Diener (2002) offer comprehensive reviews. An important early result, sometimes referred to as "well-being paradox" (Easterlin 1974, 1995), is that the average satisfaction in a country does not increase as countries grow wealthier. At the individual level, there is a weak positive cross-sectional association between income and SWB. If one follows an individual over the life-cycle however, as income first increases and then levels off, happiness remains unchanged.

Empirical explanations for the paradox have ventured in two directions. One literature focuses on the moderating role of income expectations, aspirations, as well as adaptation. Another literature – the one our paper contributes to – conjectures that a single-item SWB measure may be too restrictive to fully capture the impact of income on quality of life. In this view, there is a crucial distinction between positive and negative well-being, borrowing from psychology, where positive well-being is related to good feelings such as happiness, self-esteem, or life-satisfaction, whereas negative well-being is related to bad emotions such as depression, anxiety, or distress.

Psychologists claim a certain degree of independence between these two, since important determinants of positive well-being are found to have less influence on negative well-being, and vice versa (Huppert and Whittington 2003). The empirical strategy to identify such asymmetries is to find one measure of positive and one measure of negative well-being (e.g., life-satisfaction versus psychological stress) and to perform two separate analyses with equal regressors. In a recent paper, Headey and Wooden (2004) translate this concept to economic circumstances and show, among other things, that income and employment status have stronger effects on financial satisfaction than they have on financial stress.

We argue in this paper that the distinction between positive and negative well-being can and should be made even if one uses only a single-item measure of well-being. Roughly speaking, movements in the left part of the outcome distribution represent more or less dissatisfaction (or unhappiness), whereas movements in the right part of the outcome distribution represent more or less satisfaction (or happiness). The model we propose allows income to affect satisfaction and dissatisfaction in different ways, relative to the benchmark models used in the previous literature. Specifically, we develop a panel data version of the Maddala/Terza varying threshold ordered response model (Maddala 1983, Terza 1985). The model is based on multiple indices, possibly but not necessarily varying by the level of SWB, and it allows for unobservable individual specific effects.

With data from the German Socio-Economic Panel we find that income significantly reduces negative SWB but it does not affect positive SWB in a subsample of men living in one-person households. This finding corroborates previous evidence of asymmetric effects from multi-item analyses, this time with a single-item measure. An immediate implication of response asymmetries is a variable compensating income variation. For example, the implied income compensation for unemployment differs by a factor of six compared to the standard models, depending on the position of the individual on the SWB scale. For women, the evidence for the asymmetry hypothesis is less clear cut.

The rest of this paper is structured as follows. In the next section, we offer a brief review of the literature on income and SWB. One important lesson is that using a single measure of SWB together with a sufficiently flexible econometric model has clear advantages for the calculation of compensating income variation, as it is increasingly applied in the valuation of public goods and intangibles. The econometric models are presented in Section 3. The data and the estimation results are discussed in Sections 4 and 5, respectively. Section 6 concludes.

2 Background

Empirical evidence on the relationship between income and subjective well-being (SWB) is important for (at least) two reasons. First, the design and evaluation of economic policies often takes income as the target quantity of interest. The idea is, of course, that income is a good proxy for well-being, and that it is easy to measure. If the link between income and well-being is less strong than suspected, then economic policies based on income (or GDP) maximization alone may turn out to be inferior from an overall well-being perspective.

Second, the relationship between income and well-being may be used to put a monetary value – or shadow price – on non-traded goods, usually in the context of cost-benefit analyses. The basic idea is one of compensation: in case of a "bad", how much of an increase in income is required to offset the negative effect of the bad, while keeping the person at the same level of SWB as in the absence of the bad? Similarly, in case of a good, one can implicitly determine the shadow price of the good by asking how much income a person would be willing to give up in order to obtain the good. Examples for this line of research are Blanchflower and Oswald (2004), who estimate the pecuniary value of a lasting marriage (relative to widowhood) to be $100,000 per year. Other examples include Winkelmann and Winkelmann (1998) who estimate the money-equivalent value of the psychological cost of unemployment, and Schwarze (2003) who uses the principle to determine an income equivalence scale, i.e., the income compensation required to keep the same level of an individual's well-being with one additional household member present. Frey, Luechinger and Stutzer (2004) estimate the value of public safety, or the absence of terrorism. Van Praag and Baarsma (2005) measure the external cost of air traffic noise for people living near the Amsterdam Airport.

Unfortunately, the implied compensation may be sensitive to the chosen model, and too restrictive assumptions may lead to spurious estimates. Usually, income is entered in logarithmic form, either in a linear regression framework or in a conditional probability model such as the ordered probit model. In this case, the marginal effects are inversely proportional to income: to achieve the same increase in (average) happiness, larger and larger absolute changes in income are necessary. While semi-parametric estimators usually provide some support for a log-linear functional form, they still preclude the occurrence of asymmetric effects as defined in the previous section.

Only recently have systematic attempts been made to allow for more flexible response patterns by using generalized models. The heart of the matter is the single index structure of the prevailing models. In single-index models the dependent variable, say y, is modeled as a function of a linear combination of independent variables, say x. For example, the linear regression model is a single-index model with $E(y|x) = x'\beta$ with parameter vector β. Other examples include the probit model or the ordered probit model. Contrary to that, in multiple-index models, such as the multinomial logit model, the relationship between y and x depends on more than one linear combination of x, i.e., on $x'\beta_1$, $x'\beta_2$, etc. The key to more flexible models of the income/well-being relationship lies in the formulation of multiple-index models that respect the ordered nature of the dependent variable.

A recent example for such an approach is the latent class analysis presented in Clark et al. (2005). They identify four distinct classes (and hence four separate index functions), using data from the European Community Household Survey. In their application, the marginal probability effects of an income change are larger in the "latent happy" than in the "latent unhappy" classes.

We propose in this paper an alternative approach with outcome-specific parameters. The technical details are discussed in the next section. The important point to notice is the close relationship between our approach and the psychological distinction between positive and negative well-being (e.g., Diener 1984, Diener and Emmons 1984, Headey et al. 1993, Diener et al. 1999). While this literature models asymmetries in the well-being response by using two separate measures of the left hand side variable, one of well-being (e.g., life-satisfaction or financial satisfaction) and the other of ill-being (e.g., mental health or financial stress), we allow for such asymmetric effects in a single-item well-being measure.

The potential benefit of the suggested approach is that compensating variations can be calculated without taking into account a complex correlation structure between positive and negative measures of well-being. Moreover, we do not rely on the availability of such measures, instead we can directly access large databases containing only one measure of well-being, such as life-satisfaction. Our approach is closely related to the flexible estimation of marginal probability effects, i.e., what we are looking at is how much the probability of each outcome in the SWB distribution responds to a marginal increase in income. Compared to the previous literature, we propose a new panel data model that is very flexible in the way income affects SWB and additionally allows for individual specific effects.

3 Econometric Modeling

Most empirical work on the determinants of SWB so far uses either linear regression or single-index ordered response models. While the latter account for the discreteness and ordering of the dependent variable, they impose an implicit cardinalization such that, for example, the trade-off ratios between income and other determinants of well-being must be constant across the distribution of outcomes (Winkelmann and Boes 2006). Since we want to estimate unrestricted income effects for low and high levels of well-being, we need to use a more flexible model, and we propose a generalization of Maddala's (1983) and Terza's (1985) model to panel data.

Let $y_{it} \in \{1, \ldots, J\}$ denote SWB of individual $i = 1, \ldots, n$ at time $t = 1, \ldots, T_i$ obtained from the survey response to the general satisfaction question, and let x_{it} denote a vector of (possibly time-varying) covariates. The relationship between y_{it} and x_{it} can be specified in terms of cumulative conditional probabilities:

$$P(y_{it} \leq j | x_{it}; \theta_j) = \Phi(-x_{it}\theta_j) \qquad j = 1, \ldots, J-1 \tag{1}$$

where $\Phi(\cdot)$ denotes the cumulative density function of the standard normal distribution, and θ_j denotes a vector of category-specific parameters including a constant term. In order to ensure positive cell probabilities, we require the θ_j's to fulfill the strict inequalities $x_{it}\theta_1 > \ldots > x_{it}\theta_{J-1}$. Rewriting $x_{it}\theta_j$ as

$$x_{it}\theta_j = \alpha_j + \tilde{x}_{it}\beta_j \qquad j = 1, \ldots, J-1 \tag{2}$$

we see that the standard ordered probit model is nested for $\beta_1 = \ldots = \beta_{J-1}$, i.e., under the hypothesis of equal slope parameters. The cumulative probabilities in (1) form a well-defined conditional probability model which can be estimated by standard maximum likelihood methods to obtain consistent and asymptotically normal estimates of θ. In order to provide valid inference, the standard errors should be adjusted for clustering at the individual level.

In order to exploit the advantages of panel data more fully, the model can be augmented by individual specific time invariant effects. Conditioning on such effects avoids bias if, for example, unobserved personality traits affect SWB as well as observable characteristics (see Ferrer-i-Carbonell and Frijters 2004). Let c_{ij} denote individual effects that are constant over time but possible vary by the SWB level, and rewrite (1) conditional on c_{ij}

$$P(y_{it} \leq j | x_{it}, c_{ij}; \theta_j) = \Phi(-x_{it}\theta_j - c_{ij}) \qquad j = 1, \ldots, J-1 \tag{3}$$

or in terms of a conditional probability model

$$\begin{aligned} P(y_{it} = 1 | x_{it}, c_i; \theta) &= \Phi(-x_{it}\theta_1 - c_{i1}) \\ P(y_{it} = j | x_{it}, c_i; \theta) &= \Phi(-x_{it}\theta_j - c_{ij}) - \Phi(-x_{it}\theta_{j-1} - c_{ij-1}) \qquad j = 2, \ldots, J-1 \\ P(y_{it} = J | x_{it}, c_i; \theta) &= 1 - \Phi(-x_{it}\theta_{J-1} - c_{iJ-1}) \end{aligned} \tag{4}$$

where $\theta = (\theta_1' \ \ldots \ \theta_{J-1}')'$ and $c_i = (c_{i1} \ \ldots \ c_{iJ-1})'$. We assume that x_{it} is strictly exogenous conditional on c_i and that outcomes are independent conditional on (x_i, c_i), where x_i contains x_{it} for all t. The first assumption rules out lagged dependent variables in x_{it}, the second assumption allows for dependencies in y_{it} across t if conditioned only on observables x_i.

Without specifying the relationship between x_{it} and c_i, i.e., treating c_i as fixed parameters to be estimated along with θ, model (4) introduces an incidental parameters problem. For fixed time and large cross-sectional dimension, the number of parameters c_i is unbounded, with available information on c_i being fixed, which yields inconsistent estimators of c_i and θ. We solve this problem by treating c_i as random variable drawn along with (x_i, y_i) and following the idea of Chamberlain (1980) under a Mundlak (1978) restriction to allow for possible correlation between c_i and x_i:

$$c_{ij} = \bar{x}_i \gamma_j + \alpha_i \tag{5}$$

where $\bar{x}_i$ is the average of x_{it} over time, γ_j is a conformable parameter vector, and α_i is an orthogonal error with $\alpha_i | x_i \sim Normal(0, \sigma_\alpha^2)$. The joint distribution of $(y_{i1}, \ldots, y_{iT_i})$ conditional on observables is then obtained by integrating out α_i in the probabilities (4),

$$f(y_{i1}, \ldots, y_{iT_i} | x_i; \theta, \gamma, \sigma_\alpha) = \int_{-\infty}^{\infty} \prod_{t=1}^{T_i} \prod_{j=1}^{J} P(y_{it} = j | x_{it}, \bar{x}_i, \alpha_i; \theta, \gamma)^{\mathbf{1}(y_{it}=j)} \frac{1}{\sigma_\alpha} \phi\left(\frac{\alpha_i}{\sigma_\alpha}\right) d\alpha_i \tag{6}$$

with $\gamma = (\gamma_1' \ \ldots \ \gamma_{J-1}')'$ and indicator function $\mathbf{1}(\cdot)$. The integral in (6) does not have a closed form solution, but with a simple change of variables from α_i to $\psi_i = \alpha_i / (\sqrt{2}\sigma_\alpha)$ it can be rewritten in a form amenable to Gauss-Hermite quadrature for numerical approximation. Estimation of parameters by maximum likelihood is straightforward once the integral has been evaluated, and the resulting estimator is consistent, efficient and approximately normally distributed with covariance matrix equal to the inverse of the expected Hessian. Model (3) with random effects specification has been implemented in a new Stata module called *regoprob* available on the author's homepage *www.unizh.ch/sts/* or via the *ssc* commands in Stata.

4 Data

The German Socio-Economic Panel (GSOEP) is an annual panel survey of randomly selected households in Germany (see Burkhauser et al. 2001). Personal information is available for all household members aged 16 and above. Our data are drawn from the West Germans (A) subsample 1984-2004, yielding a maximum of 21 observations per individual. We apply a number of standard selection criteria: included individuals are between 25 and 65 years old at the time of the survey, and we require non-missing information on all the included variables.

In addition, we employ a novel restriction by considering single person households only. The rationale for this selection is that the match between reported household income and individual material well-being is much better in single-person households than we could possibly hope for in a multi-person household. General household surveys such as the GSOEP typically include two types of income measures, one being total household income (from all sources), the other being personal labor earnings. Clearly, personal labor earnings are not a very good indicator of material well-being, in particular, but not only for persons who do not work, as it does not include any government transfers (e.g., child benefit, government grants, or rent subsidies). Household income (net of taxes and social security contributions) is in general a more appropriate measure. However, in multi-person households, there remain two types of ambiguities. First, there is an ongoing debate on the right equivalence scale in order to reflect economies of scale in household production and consumption. Secondly, we do not know whether resources are shared evenly within the household, but such an (arbitrary) assumption is required when assigning one income to several household members.

For these reasons, we find it instructive to study the relationship between income and SWB in the context of single person households. Of course, this raises the question of external validity: to what extent can results for single person households be extrapolated to the population of all households? While single person households are clearly selected on a number of factors (such as age, and possibly also income) it is a-priori unclear why the relationship between income and SWB should be different for such persons.

All in all, this approach leaves us with 5007 person-year observations for men, and with 4727 person-year observations for women. The dependent variable is, as mentioned before, the response to the survey question "How satisfied are you with your life, all things considered?". There are relatively few responses in the 0-2 range. For this reason, and to preserve some degrees of freedom (a full set of regression parameters is added for each additional category), we use a modified 0-9 scale, where the original 0-2 responses have been grouped into the lowest "dissatisfied" category.

Figure 1 depicts the frequency distribution of SWB responses in our sample, separately for men and women. People are mostly satisfied with their life: about two thirds report a SWB level of seven or higher, and women have a slightly higher average SWB level than men. The distribution in Figure 1 is characteristic of most SWB distributions in the sense that the majority of people responds a relatively high level of SWB, although the highest response category is chosen relatively infrequently.

— Insert Figure 1 about here —

In the regression analysis, control variables include - apart from income - a second order polynomial in age and dummy variables for unemployment and health status. The health dummy indicates good health status and has been generated using the satisfaction with health response. Satisfaction with health is originally measured on the eleven-point scale (0-10) and we dichotomized at the median value of seven. Of course, household status and gender is controlled for as well by the way the sample is set up. We use a relatively simple specification with only a few variables. This has two main advantages. First, since eight regression parameters are estimated for each variable, fewer regressors keep the model manageable. Second, many of additional variables used in the previous literature are arguably endogenous choice variables, obstructing the interpretation of the results.

Table 1 summarizes the sample means of the explanatory variables by gender. Among one-person households, men have a significantly higher monthly income than women (about 260 Euros) and are on average more than five years younger. The unemployment rate is about 2.5 percentage points higher for men than for women, and 58.2 percent of the women are relatively satisfied with their health status (compared to 65.6 percent of the men). These variations can largely be explained by the different age distributions of single male and single female households. Men are mostly living alone when they are young and at the beginning of their career path. Women are more likely to live alone when they are older, contributing factors being a higher incidence of widowhood due to greater life-expectancy.

— Insert Table 1 about here —

Table 2 cross-tabulates the sample means of the dependent variable conditional on the SWB response, again separately for men and women. The income variable shows a lot of variation along the SWB dimension. For men (panel A), the lowest average monthly income (1124 Euro) is observed for individuals with very low SWB, the highest income (1519 Euro) for those with response "8". When moving from the utmost left part of the SWB distribution to the right, average income is first increasing then decreasing. A similar pattern can be observed for women (panel B), although on a lower level. Concerning the rest of the variables, a common result known from the previous literature and also confirmed by the present data is that among less satisfied people the unemployment rate is relatively high, and that reported health status and SWB are positively correlated.

— Insert Table 2 about here —

5 Estimation Results

In this section, we report on the estimation results of the relationship between income and subjective well-being. We first present the estimated income parameters under several model assumptions, then turn our attention to the implications with respect to the asymmetry hypothesis, and finally discuss the robustness of our results.

Two different models of subjective well-being have been estimated. The first is a standard ordered probit model (OProbit). The second is the generalized ordered probit model (GOProbit), where all parameters are outcome-specific. In order to check for model robustness, we also report the results of a series of eight random effects binary probit models (DProbit), where the dependent variables result from dichotomization of the SWB response, i.e., $y_{it} > 2$ against $y_{it} \leq 2$, $y_{it} > 3$ against $y_{it} \leq 3$, and so on. The binary regressions consistently estimate the parameters θ in the cumulative probabilities (3) without imposing the ordering. By contrast, the generalized ordered probit model takes into account the ordering and is therefore more efficient.

Table 3 displays the estimated income coefficients separately for men (panel A) and women (panel B). For men, we find a positive and significant parameter in the standard model (0.362 with z-value 6.67) indicating a positive relationship between income and SWB. In the generalized model, eight different parameters are estimated. The income coefficients are slightly higher for the parameter vectors θ_1 to θ_6 than the overall estimate in the standard model. The point estimate decreases but is still significant for θ_7, and finally turns negative and insignificant for θ_8. When comparing the generalized ordered and the binary models we notice very similar results.

— Insert Table 3 about here —

The estimated coefficients in the sample of women are smaller (in absolute value) and less significant than those for men indicating a weaker relationship between income and SWB. For example, in the standard model we obtain a point estimate of 0.131, which is only about a third of that for men, and the z-value decreases to 1.97. Interestingly, in the generalized model the income coefficients are significant on the 5%-level only for θ_4 and θ_5, while all other income coefficients are insignificant. As for men, the results from the binary regressions support the generalized model. If we formally test the generalized ordered probit model against the standard model, we can reject the null hypothesis of equal slope parameters for men ($LR_{203} = 548.9$) and for women ($LR_{203} = 430.1$). The null hypothesis of equal income coefficients is also rejected in both cases.

There are a number of ways to interpret the estimated parameters, but we focus here on two concepts that offer a very intuitive interpretation when dealing with conditional probability models. The first is the marginal probability effect (MPE) of income and the second the trade-off ratio between income and unemployment. The former shows how marginal changes in income affect the distribution of SWB responses, which is of particular interest for the asymmetry hypothesis since we are able to identify whether the effect of income on SWB differs for low and high SWB. The latter concept provides an extension to the discussion of asymmetric effects by answering the question of how much income is necessary to compensate for unemployment, given the levels of SWB. It differs from the first, as it implicitly also responds to asymmetric effects of unemployment throughout the distribution of outcomes.

MPE's can be obtained by taking first derivatives of (4) with respect to the variable(s) of interest. However, since the c_{ij}'s are random variables, we cannot directly calculate the MPE's without further assumptions. One possibility would be to set c_i equal to its expected value and proceed with standard calculus. Alternatively, due to the normality assumptions, the cumulative probabilities marginal on c_i are given by (1), where θ needs to be replaced by $\vartheta = \theta(1 + \sigma_c^2)^{-1/2}$. This is called the populationaveraged coefficient vector. It follows that

$$\begin{aligned} MPE_1^{(l)} &= \frac{\partial P(y_{it}=1|x_{it};\vartheta)}{\partial x_{it}^{(l)}} = -\phi(-x_{it}\vartheta_1)\vartheta_1^{(l)} \\ MPE_j^{(l)} &= \frac{\partial P(y_{it}=j|x_{it};\vartheta)}{\partial x_{it}^{(l)}} = \phi(-x_{it}\vartheta_{j-1})\vartheta_{j-1}^{(l)} - \phi(-x_{it}\vartheta_j)\vartheta_j^{(l)} \qquad j=2,\dots J-1 \\ MPE_J^{(l)} &= \frac{\partial P(y_{it}=J|x_{it};\vartheta)}{\partial x_{it}^{(l)}} = \phi(-x_{it}\vartheta_{J-1})\vartheta_{J-1}^{(l)} \end{aligned} \tag{7}$$

where $x_{it}^{(l)}$ denotes the l-th element in x_{it} (here assumed to be income) and $\vartheta_j^{(l)}$ the corresponding scaled (income) coefficient. Discrete probability changes can be approximated by $\Delta P(y_{it}=j|x_{it};\vartheta) \approx MPE_j^{(l)}\Delta x_{it}^{(l)}$. Obviously, the MPE's are functions of x_{it} and therefore differ for each individual. We evaluate the MPE's at the sample averages of the regressors.

Table 4 summarizes the marginal probability effects of income by gender. Consider, for example, the results for men and take the *ceteris paribus* effect of increasing logarithmic household income by a small amount on the probability of responding a SWB level of "8". Table 4 shows a value of 0.059 for the standard model. This means that the probability of a response of "8" increases by 0.059 percentage points if we increase logarithmic income by 0.01, which corresponds approximately to a one-percent increase in level income. A doubling of income, i.e., a change in logarithmic income by 0.693, increases the probability of response "8" by about $0.059 \times 0.693 \times 100$, or about 4.09 percentage points, ceteris paribus.

— Insert Table 4 about here —

Comparing the MPE's among the three different models and over all possible outcomes, we obtain the following main results. For men all models suggest that more income significantly reduces the probability of low SWB (0-5), and significantly increases the probability of response "8". For high SWB responses (9-10), the standard model predicts a strong positive relationship between income and SWB, whereas the generalized model and also the binary models do not find a significant effect. Since the restricted OProbit is clearly rejected, we conclude that income has no effect on positive well-being. Our preferred specification supports the asymmetry hypothesis for men: higher income decreases the probability of negative well-being (low SWB), but it does not affect the probability of positive well-being (high SWB).

For women the relationship between income and SWB is relatively weak. While the standard model finds small but significant effects for low and high SWB responses, the generalized model predicts a significant negative effect only on the probability of responses "5" and "6". The gender difference might be explained by social norms that assign the role of primary income earner to men and therefore make income a relatively more important determinant of male well-being (see also Lalive and Stutzer 2004).

An alternative interpretation of the coefficients assesses the importance of income *relative* to other determinants. Take unemployment as an example: what is the income increase required to offset the negative well-being effect of unemployment? While unemployment is a binary variable, marginal changes in unemployment, or in this case more precisely, the sample unemployment rate, are meaningful as long as we are interested in average MPE's. It follows from totally differentiating that

$$dP(y_{it} = j|x_{it}; \vartheta) = MPE_j^{(l)} dx_{it}^{(l)} + MPE_j^{(m)} dx_{it}^{(m)} \qquad j = 1, \ldots, J \tag{8}$$

where $x_{it}^{(l)}$ and $x_{it}^{(m)}$ denote income and unemployment, respectively, and the MPE's are given by (7). The approximation in (8) directly leads to the concept of compensating variation: How much of a variation in one regressor (here income) is needed to offset the given change in another regressor (here unemployment) such that $dP(y_{it} = j|x_{it}; \vartheta) = 0 \; \forall j$, i.e., all probabilities remain unchanged. Rearranging terms yields

$$\frac{dx_{it}^{(l)}}{dx_{it}^{(m)}} = -\frac{MPE_j^{(m)}}{MPE_j^{(l)}} \qquad j = 1, \ldots, J \tag{9}$$

In the standard model, this trade-off ratio reduces to the ratio of coefficients which does not vary across outcomes, whereas in the generalized model such an restriction is not imposed.

Table 5 shows the required changes in logarithmic income if the (sample) unemployment rate increases by one percentage point, $dx_{it}^{(m)} = 0.01$, given the SWB distribution is fixed. If we want to interpret the reported numbers, we need to be careful with respect to the significance of MPE's. The trade-off ratio does only make sense for significant income effects. In this case, the required change in income is either zero if the MPE of unemployment is statistically not different from zero, or the change is positive (or negative) for significant unemployment effects. We marked the four cases (non-sensible/zero/positive/negative) with $\times / \circ / + / -$.

— Insert Table 5 about here —

The numbers in Table 5 (multiplied by 100) approximate the percentage change in income, e.g., for men in the standard model a 0.019 means that income must increase by 1.9 percent to offset the increase in the unemployment rate by one percentage point. As expected, the trade-off ratios in the ordered probit model are constant for all levels of SWB, and interpretation therefore is not particularly interesting. In the generalized model, required income changes vary between 0.6 and 4.2 percent. An important observation is that income compensations are entirely ineffective for men with high SWB, and effective for medium to low satisfied men, though in a very heterogeneous way. For women, a compensation for unemployment in terms of income is rather unpromising, and other factors determining SWB need to be identified when looking for effective compensation schemes.

While these results are obtained for a specific sample and a specific parametric model with its set of assumptions, we found a remarkable robustness of the main conclusions with respect to alternative specifications and samples. Possible alternatives include the use of different link functions (rather than the probit ones), including the logit, the log-logistic, and the complementary log-log; conditioning on fixed effects using Chamberlain's (1982) conditional logit model; the use of multi-person household samples; and possible endogeneity of income in the SWB equation. We could not find evidence for endogeneity. Neither was the probit assumption rejected, nor did the response asymmetry for men disappear in multi-person households.

6 Conclusion

This paper reconsiders the relationship between income and subjective well-being. Recent empirical evidence suggests that income might have asymmetric effects on positive and negative well-being. In this literature, the empirical strategy to identify such asymmetries is to use two separate measures of well-being, e.g., life-satisfaction and psychological stress. This approach has many disadvantages compared to the usage of a single measure, e.g., the lack of availability of such multi-dimensional data, or the problem of possibly complex correlation structures when one is interested in compensating variations. We therefore propose a new and very flexible panel data model with which we can analyze whether income effects depend on the level of well-being. The model allows for individual specific effects and outcome-specific parameters, the latter feature being particularly important for income to affect SWB in a non-monotonic way, i.e., the effects depend on the SWB level.

In a sample of men in single-person households drawn from the German Socio-Economic Panel waves 1984 to 2004, we find strong support for the hypothesis of asymmetric income effects: while income significantly reduces the probability of low SWB responses, it does not affect high SWB. For women in single-person households income plays a minor role in the formation of SWB, and support of the asymmetry hypothesis is rather weak. In an extension to cost-benefit considerations, we calculate the trade-off ratio between income and unemployment and show that the required income compensations vary between 0.6 and 4.2 percent. These are far from being constant – a restriction that is imposed in the standard ordered probit model.

Thus, by investigating marginal probability effects and trade-off ratios we were able to document an asymmetric impact of income on SWB using a single measure of well-being, namely general life-satisfaction. If one wants to estimate marginal probability effects and compensating variations in a meaningful way, then one should use the generalized ordered probit model rather than the simpler models prevailing in earlier research.

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Figure 1: Marginal Distribution of Subjective Well-Being Responses

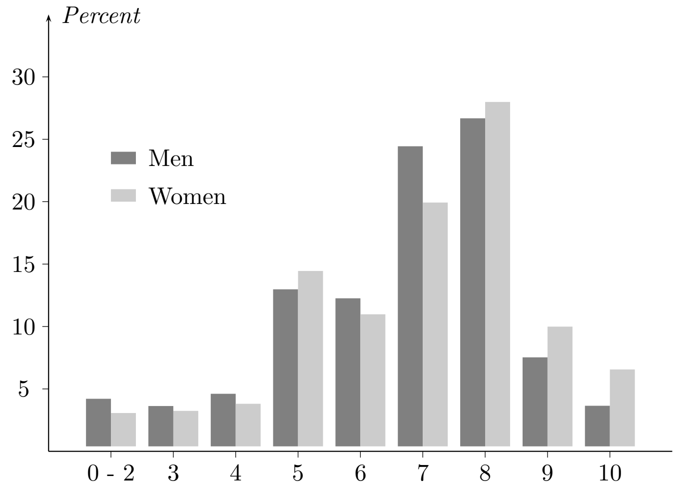

Table 1: Descriptive Statistics by Sex

	Men		Women	
Variable	Mean	Std.Err.	Mean	Std.Err.
Monthly income in EUR	1403.5	12.0	1140.9	10.3
Age in years	40.24	0.16	45.80	0.20
Unemployment (0/1)	0.083	0.004	0.058	0.003
Good health (0/1)	0.656	0.007	0.582	0.007
Number of Obs.	5008		4727	

Table 2: Sample Means by Sex and Well-Being Level

	Subjective Well-Being								
Variable	0 - 2	3	4	5	6	7	8	9	10
A. Men									
Relative Freq.	4.21%	3.63%	4.61%	12.98%	12.26%	24.44%	26.68%	7.53%	3.65%
Income	1123.9	1152.0	1414.3	1255.8	1324.0	1477.9	1519.3	1473.6	1267.8
Age	43.86	41.83	41.07	43.55	40.32	38.99	38.91	38.19	43.79
Unemployment	0.336	0.176	0.182	0.126	0.103	0.056	0.030	0.029	0.022
Good Health	0.336	0.319	0.338	0.340	0.549	0.732	0.841	0.897	0.896
B. Women									
Relative Freq.	3.07%	3.24%	3.81%	14.45%	10.98%	19.93%	27.99%	9.99%	6.56%
Income	930.5	935.4	1047.4	978.1	1055.7	1196.7	1238.6	1290.7	1082.1
Age	47.50	45.75	45.59	49.28	46.25	43.37	44.62	45.34	49.84
Unemployment	0.234	0.124	0.172	0.089	0.052	0.036	0.039	0.013	0.035
Good Health	0.159	0.196	0.267	0.274	0.420	0.601	0.767	0.847	0.845

Table 3: Estimated Income Coefficients by Sex and Well-Being Level

		Parameter vector							
	overall	θ_1	θ_2	θ_3	θ_4	θ_5	θ_6	θ_7	θ_8
A. Men									
OProbit	0.362								
	(0.054)								
GOProbit		0.417	0.470	0.423	0.372	0.429	0.456	0.314	-0.066
		(0.141)	(0.109)	(0.093)	(0.082)	(0.079)	(0.082)	(0.115)	(0.152)
DProbit		0.396	0.412	0.473	0.351	0.442	0.413	0.228	-0.102
		(0.153)	(0.118)	(0.103)	(0.086)	(0.082)	(0.082)	(0.118)	(0.183)
B. Women									
OProbit	0.131								
	(0.067)								
GOProbit		0.256	0.060	0.144	0.234	0.329	0.175	-0.070	-0.045
		(0.225)	(0.144)	(0.131)	(0.108)	(0.100)	(0.098)	(0.117)	(0.146)
DProbit		0.039	0.014	0.128	0.196	0.311	0.162	-0.066	0.012
		(0.199)	(0.157)	(0.128)	(0.107)	(0.101)	(0.097)	(0.120)	(0.169)

Notes: The models are the standard ordered probit (OProbit), the generalized ordered probit (GOProbit), and a series of binary probit models (DProbit), where the dependent variables result from dichotomization of the SWB response. The estimates are reported for logarithmic income. Each model controls for a quadratic form in age, unemployment (0/1), good health (0/1), and time fixed effects. Individual effects are assumed to be decomposable into a linear function of individual group means and orthogonal error, and the likelihood for each individual is approximated using Gauss-Hermite quadrature.

Table 4: Marginal Probability Effects of Income by Sex and Well-Being Level

	Subjective Well-Being								
	0 - 2	3	4	5	6	7	8	9	10
A. Men									
OProbit	-0.016	-0.014	-0.016	-0.037	-0.020	0.003	0.059	0.027	0.014
	(0.003)	(0.001)	(0.001)	(0.003)	(0.009)	(0.003)	(0.009)	(0.005)	(0.005)
GOProbit	-0.020	-0.022	-0.014	-0.027	-0.037	-0.005	0.088	0.039	-0.002
	(0.007)	(0.006)	(0.004)	(0.005)	(0.006)	(0.007)	(0.033)	(0.109)	(0.089)
DProbit	-0.019	-0.017	-0.029	-0.018	-0.048	-0.011	0.094	0.028	-0.003
	(0.006)	(0.008)	(0.010)	(0.017)	(0.019)	(0.023)	(0.022)	(0.011)	(0.006)
B. Women									
OProbit	-0.004	-0.005	-0.005	-0.016	-0.008	-0.003	0.020	0.012	0.008
	(0.002)	(0.001)	(0.001)	(0.005)	(0.012)	(0.003)	(0.011)	(0.004)	(0.006)
GOProbit	-0.009	0.005	-0.011	-0.036	-0.040	0.038	0.064	-0.008	-0.003
	(0.008)	(0.016)	(0.020)	(0.015)	(0.013)	(0.029)	(0.116)	(0.125)	(0.027)
DProbit	-0.001	0.001	-0.011	-0.031	-0.040	0.034	0.059	-0.011	0.001
	(0.006)	(0.007)	(0.013)	(0.018)	(0.021)	(0.026)	(0.028)	(0.020)	(0.012)

Notes: The models are the standard ordered probit (OProbit), the generalized ordered probit (GOProbit), and a series of binary probit models (DProbit). The marginal probability effects have been calculated for logarithmic income evaluated at the sample means of the explanatory variables and marginal on the individual effect. For the binary models, we approximated the cross-parameter covariances by the method proposed by White (1982). An increase in income by one percent corresponds to an increase in logarithmic income by 0.01, i.e., reported numbers can be interpreted directly as change in percentage points. Standard errors are approximated using the delta method.

Table 5: Trade-Off Ratios Between Income and Unemployment

	Subjective Well-Being								
	0 - 2	3	4	5	6	7	8	9	10
A. Men									
OProbit	0.019^{+}	0.019^{+}	0.019^{+}	0.019^{+}	0.019^{+}	$0.019^{\times}$	0.019^{+}	0.019^{+}	0.019^{+}
GOProbit	0.009^{+}	0.006^{+}	0.042^{+}	0.015^{+}	0.010^{+}	$0.028^{\times}$	0.011^{+}	$0.011^{\times}$	$-0.179^{\times}$
B. Women									
OProbit	$0.026^{\times}$	0.026^{+}	0.026^{+}	0.026^{+}	$0.026^{\times}$	$0.026^{\times}$	$0.026^{\times}$	$0.026^{\times}$	$0.026^{\times}$
GOProbit	$0.035^{\times}$	$-0.029^{\times}$	$0.029^{\times}$	0.007^{+}	-0.004°	$0.024^{\times}$	$-0.008^{\times}$	$-0.094^{\times}$	$0.102^{\times}$

Notes: The models are the standard ordered probit (OProbit) and the generalized ordered probit (GOProbit). The trade-off ratios show the required change in logarithmic income to compensate for an increase in the (sample) unemployment rate by one percentage point, given the probability of a certain SWB response remains unchanged. The ratio of significant marginal income and unemployment effects is marked $+/-$ (positive/negative). If the marginal income effect is insignificant, the ratio is marked $\times$. If the income effect is significant but the unemployment effect is not, the ratio is marked $\circ$.

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